Multimarket contact, competition and pricing in banking

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ABSTRACT

In this paper we propose an empirical test of the multimarket contact theory for the banking industry by assessing the impact of multimarket linkages of banks on their market power. This is done by means of a simultaneous equation model, in which the divergence of price from marginal cost is considered as a function of multimarket contact. The model is estimated using aggregate data from the Italian regions for the years 1997-2009. Our results show that multimarket linkages among banks are positively and significantly correlated to the market power index, and are also connected to both home and non-home market concentration. Their effect on loan price appears not negligible. The evidence is robust to changes in model specification and multimarket contact measures, and supports the idea that firms experimenting a larger number of contact across multiple markets are more likely to collude. This finding may have antitrust implications, considering that the latest trend towards consolidation in the banking sector has notably increased the geographic overlap of banks' branch networks and thus the multimarket linkages among them.

KEYWORDS: Banking; Competition; Market structure; Conduct; Regional analysis

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1. Introduction

Firms face multimarket competition when they compete with each other at the same time in several different markets. The economic literature has generally considered multimarket contact as a device for lessening competition and/or facilitating collusion in a market. Since Edwards (1955), the idea is that, when one large firm competes with another, they are likely to encounter each other in a considerable number of markets, and the multiplicity of contact may dampen the edge of their competition. Accordingly, the "multimarket contact theory" – also known as "mutual forbearance theory" or the "linked oligopoly theory" – maintains that firms operating in the same (geographical or product) markets may have less incentive to compete aggressively in a given market if they fear rivals' retaliation in all the other markets.

In this paper we test this hypothesis within a two-equation model of competition for the Italian banking industry in the years 1997-2009. Particularly, we develop a simultaneous model with a demand equation and an optimality equilibrium condition (i.e. marginal revenue equal to marginal cost) where a parameter measuring the average degree of market power of banks is included. We suppose that this parameter is a function of multimarket contact (and also of market concentration) – a novelty within the analysis of banking industries, to the best of our knowledge – and estimate the model using panel data from the twenty Italian regions.

The reasons for the focus on a banking market are twofold. First, banks can be regarded as firms selling a rather homogeneous product in multiple local markets; hence, there are potential geographical overlaps that make the banking industry an ideal testing ground for the empirical appraisal of this theory.

Second, multimarket contact among banks has relevant policy implications. In the last decades, a noticeable consolidation trend has characterized the banking markets of many countries, stemming from a number of factors, like technological progress, globalization, and deregulation.[1] This phenomenon might have had significant consequences in terms of change in the degree of competitiveness, because in more concentrated markets it could be easier for firms to collude and

[1] An exhaustive analysis of aspects influencing the consolidation of the financial services industry is provided by Berger et al. (1999). In Europe, a major role in the wave of banks' mergers and acquisitions has been played by the implementation of the Economic and Monetary Union (EMU).

therefore set higher prices. Besides, mergers and acquisitions (M&A) have led to a larger overlap of branch networks, thus increasing the number of markets where banks meet. So, if the multimarket contact hypothesis holds, the geographical branch structure of those banks that are involved in M&A operations becomes a key element to be considered by the antitrust authorities (Coccorese and Pellecchia, 2009, p. 246).

Similar to other European countries, the Italian banking industry has undergone both deregulation and consolidation. Since 1990, when the Central Bank of Italy removed the restrictions to branch opening, the number of banks has dropped from 1156 to 788 in 2009 (-31.8%), while the number of branches has increased from 17721 to 34036 in 2009 (+92.1%). Hence, nowadays less banks operate with more branches, so many of them meet in more markets than before. This element plainly justifies our testing of the multimarket contact hypothesis for the particular case.

The Italian banking sector is an important part of the European financial market, and shares many characteristics with a number of EU countries (e.g. France, Germany): it is a bank-oriented financial system, with rather rigid labor markets that might impede thorough restructuring and a mix of large and small banks, where a gradual shift from the traditional intermediation business to a more services-oriented industry is occurring.[2] In this respect, our analysis may represent a benchmark for similar investigations in other European nations.

The main result we get is a positive and significant link between multimarket contact and the market power index, which is also robust with respect to model specifications and multimarket contact measures. Thus, we find support to the theory according to which multimarket linkages reduce competition. A major role is also played by market concentration.

In what follows, Section 2 surveys the studies on the assessment of the level of market competition as well as on the theory and empirics of multimarket contact, with a specific focus on banking. Section 3 outlines the theoretical model, and Section 4 shows how our multimarket contact measures are calculated. Section 5 is devoted to the econometric specification of the model, while Section 6 discusses the empirical results. Section 7 concludes.

2. Market competition and multimarket contact: a review

Measuring the degree of competitiveness in the banking industry is a major theme in the economic literature, which is still divided on whether or not banking competition promotes stability. Actually, according to the “competition-fragility view” rivalry among banks triggers a reduction of

[2] See Focarelli et al (2002), p. 1048. Sapienza (2002), pp. 336-337, emphasizes that the Italian banking industry has a number of similarities also with that of the United States.

profit margins and hence encourages banks' risk taking, while the "competition-stability view" argues that more market power, by allowing banks to charge higher loan rates, is likely to increase bank risk because customers will find it harder to repay loans, thus amplifying adverse selection and moral hazard problems (Berger et al., 2009).

Empirical studies generally employ two different techniques for measuring the degree of competition in banking industries. The first is the structural approach, which originates from the traditional "structure-conduct-performance" (SCP) paradigm and is based on the idea that in highly concentrated markets it is easier to collude, so banks should enjoy higher profits. The second is the so-called "new empirical industrial organization" (NEIO) approach, which makes use of non-structural models in order to derive a conduct parameter measuring the market power exerted by banks, and entails the estimation of theoretical models of price and output determination.

Since the SCP approach (Mason, 1939; Bain, 1951) predicts a positive link between structure and performance, the standard empirical practice has been to estimate a relationship between a concentration index and a measure of performance (in terms of profits or prices). The econometric evidence coming from several SCP studies for the banking industry does not however fully support the theoretical conjecture.[3] Actually, while Berger and Hannan (1989), Hannan and Berger (1991) and Pilloff and Rhoades (2002) provide results in line with the SCP predictions, Jackson (1992), Rhoades (1995) and Hannan (1997) find less robust evidence for the positive relationship between market concentration and banks' profitability.

One reason could be that concentration indicators (such as the concentration ratio or the Herfindahl-Hirschman index) are not suitable proxies for competitive conditions: for example, Claessens and Laeven (2004) do not find any empirical evidence supporting the inverse relationship between concentration and competition, while Berger et al. (2004) emphasize that the above concentration measures have only very weak relationships with measures of profitability when the market share of firms is also included in the regression equation. The latter evidence could be justified either by the "efficient structure" hypothesis, for which high concentration endogenously reflects the market share gains of efficient firms (Rhoades, 1985; Smirlock, 1985), or by a more general problem of endogeneity characterizing the SCP tests, in which prices, profitability and concentration are all jointly endogenous (Bresnahan, 1989).

The NEIO approach has tried to overcome the shortcomings of the SCP techniques by developing methodologies that assess firms' conduct through the direct appraisal of their behaviour, thus avoiding indirect (and sometimes ambiguous) inferences about market power based on

[3] Surveys of empirical SCP studies are provided by Gilbert (1984) and Weiss (1989).

indicators of concentration. On the other side, NEIO models need detailed information on costs and demand.

One option is the estimation of a simultaneous model of demand and supply equations that includes a conduct parameter representing the behaviour of firms (and therefore the degree of their market power). It can be interpreted as either a conjectural variation coefficient or the deviation of the perceived marginal revenue schedule of a firm in the industry from the demand schedule, and can assume different values depending on the degree of market power prevailing in the industry. Pioneered by Iwata (1974), this approach has been developed by Bresnahan (1982) and Lau (1982), and applied to the banking sector by many authors (Shaffer, 1989, 1993; Berg and Kim, 1994; Shaffer and DiSalvo, 1994; Coccorese, 2005, 2012; Neven and Roller, 1999; Toolsema, 2002; Angelini and Cetorelli; 2003; Canhoto, 2004; Uchida and Tsutsui, 2005; Rezitis, 2010).

Another possibility is to estimate the so-called *H*-statistic, first proposed by Panzar and Rosse (1987), starting from a reduced form revenue equation that includes the input prices among the regressors. The *H*-statistic is calculated as the sum of the estimated elasticities of revenues to factor prices, and represents a conduct parameter that can range from negative values (corresponding to monopoly or collusion) to one (perfect competition). However, this index only allows to discriminate among different market hypotheses, even though, under specified assumptions, it could be interpreted as a continuous measure of competition (Vesala, 1995, p. 56; Bikker and Haaf, 2002, p. 2203).

The Panzar-Rosse test has been widely applied to the banking industry of several countries (Shaffer, 1982, 2002, 2004; Nathan and Neave, 1989; Molyneux et al., 1994; Coccorese, 2004, 2009; Hondroyiannis et al., 1999; De Bandt and Davis, 2000; Bikker and Haaf, 2002; Claessens and Laeven, 2004; Gelos and Roldos, 2004; Al-Muharrami et al., 2006; Casu and Girardone, 2006; Staikouras and Koutsomanoli-Fillipaki, 2006; Bikker et al., 2007, 2011; Matthews et al., 2007), especially because it only needs firm-specific data on revenues and factor prices, although the insertion of every variable shifting demand or cost (and possibly of firm-specific or also macroeconomic control variables) is most desirable.

Within the NEIO approach, a third alternative for assessing the degree of market power in banking is to calculate the Lerner index, where the marginal cost is obtained by means of the estimation of a cost function (de Guevara et al., 2005; Oliver et al., 2006; de Guevara and Maudos, 2007; Coccorese and Pellecchia, 2010). One advantage of this methodology is to provide a bank-level measure of market power, whose evolution over time can also be easily traced.

Finally, Boone et al. (2007) and Boone (2008) suggest employing the elasticity of firms' profits with respect to their cost level as a measure of competition: a higher value of this profit elasticity would indicate more intense competition. They show that it is highly correlated with the price-cost margin, but the latter tends to misrepresent the development of competition over time in markets with few firms and a high concentration, concluding that profit elasticity is a more reliable measure of competition. An application of this competition index to the lending markets of eight OECD countries has been provided by van Leuvensteijn et al. (2011).

However, the assessment of the level of competition in the banking industry would be imperfect (or even misleading) if the analysis of local markets were not combined with the evaluation of the degree of contact outside a given market among banks competing in that specific market. Actually, for several services and products competition among banks occurs at the local level, so each national banking industry can be better regarded as composed by a considerable number of separate geographic markets. In addition, phenomena like technological progress, the worldwide deregulation of capital markets and the widespread consolidation trend (through M&A) have increased the size of banks and simultaneously multiplied the occasions of encountering each other in multiple local markets (Pilloff, 1999). For these reasons, the banking sector is exposed to the influence of multimarket contact, and the degree of competition of local markets with banks that meet frequently in other markets may be different from that of local markets whose operating banks do not have any additional contact.

In the oligopoly theory, the interest in the role and effects that multimarket contact among firms may have on the intensity of competition is deep-rooted. Edwards (1955) has been probably the first to observe that multimarket firms are less willing to behave aggressively towards rivals in one market if they fear retaliation in other markets. On the contrary, they prefer to forbear from aggressive conduct in those markets in which they enjoy a competitive advantage, as long as their competitors do the same in the other markets. The reason is that, if the number of markets where firms meet is high, the retaliation would be extensive, involving multiple attacks and therefore generating notable losses that might not be balanced by a profit-increasing aggressive action in one or more markets. In such situation the overall degree of competition will be therefore reduced.

By means of a theoretical model, Bernheim and Whinston (1990) prove that multimarket contact may relax the incentive constraints governing the implicit agreements between firms and improve their abilities to sustain collusive outcomes. Besides, they show that the multimarket contact may help cooperation only if there are asymmetries either between the markets in which firms interact or between the firms within and across the multiple markets in which they meet. A refinement of this result is offered by Spagnolo (1999), who shows that multimarket links can always facilitate collusion, no matter if there are asymmetries of any type; this outcome is due to real world imperfections (e.g. managerial risk-aversion or non-linear tax effects), which make firms' objective function strictly concave. Additional theoretical support to the possibility of implicit collusion among firms in presence of multimarket contact is provided by Matsushima (2001), Thomas and Willig (2006), and Sorenson (2007).[4]

Regarding the empirical studies, two broad measures of multimarket contact are usually employed: count measures and probabilistic measures (Singal, 1996, pp. 563-564). The first group takes into account the number of meetings among firms across markets, and is the most adopted in intra-industry studies. The second group assumes a contact as a random variable, and relies on the computation of the probability of observing a smaller number of contacts than actually observed: the higher this probability, the lower is the chance that those meetings are random events.

Within manufacturing industries, and in different ways, Scott (1982), Feinberg (1985), and Hughes and Oughton (1993) give support to the mutual forbearance hypothesis, while Strickland (1985) does not find evidence for the above conjecture. Studies analyzing the role of multimarket linkages in specific industries are those by Evans and Kessides (1994), Gimeno and Woo (1996), Jans and Rosenbaum (1996), Singal (1996), Parker and Roller (1997), Fernandez and Marin (1998), Busse (2000), Gimeno (2002), Fu (2003), Waldfogel and Wulf (2006), Coronado et al. (2008), Bilotkach (2011), and Zou et al. (2012): all of them discover a positive relationship between multimarket contact among firms and entry, profitability or prices.[5]

In the empirical investigations about the influence of multimarket contact on banking competition, contact is usually measured starting from the number of links between a given group of banks (all those operating in a market, or a subset of leading institutions) or the amount of deposits involved in those links, while the level of competition is assessed through variables like profitability, lending rates, income from loans, expenses associated to deposits, changes in market shares. For this industry, the evidence is ambiguous.

Starting from the analysis of 187 U.S. major banking markets, Heggestad and Rhoades (1978) find that multimarket meetings between dominant banks adversely affect the degree of rivalry within markets (measured by the change in deposit market shares), an evidence that supports the mutual forbearance hypothesis. The same sample is used by Rhoades and Heggestad (1985), who employ traditional performance indicators (profits and prices) instead of the rivalry measure; their results provide only partial support for the linked oligopoly theory. Mester (1987) examines the

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[4] Among the theoretical contributions endorsing the hypothesis that multimarket contact increases competition, it is worth to recall the analyses by Solomon (1970), who maintains that the interconnection among banking markets may heighten competitive interaction if interbank rivalry is intense in individual local markets throughout a given region, and Mester (1992), who shows that in a finite horizon multimarket contact can have pro-competitive effects if quantity is the strategic variable.

[5] For a more comprehensive survey on the papers studying the effect of multimarket contact on competition, on both the theoretical and the empirical ground, see Coccorese and Pellecchia (2009), pp. 246-249.

savings and loan industry in California, and finds that high concentration coupled with high contact is beneficial to consumers, hence multimarket links appear to have a pro-competitive effect.

Whalen (1996) and Pilloff (1999) focus on different U.S. banking organizations, both getting results consistent with the linked oligopoly theory, while Haveman and Nonnemaker (2000) find an inverted U-shape influence of multimarket contact on growth and entry in new markets within the California savings and loan industry, with multimarket firms growing in (and entering) those markets with moderate multimarket contact and avoiding those in which the level of multimarket contact is high. Fuentelsaz and Gomez (2006) study the Spanish savings bank market, and also find a U-inverted influence of multimarket contact on entry.

The linked oligopoly theory for the Italian banking industry has been tested only by De Bonis and Ferrando (2000) and Coccorese and Pellecchia (2009). While the evidence of De Bonis and Ferrando (2000) is that geographical overlap in banking is positively correlated with changes in market shares and lower lending rates, thus refuting the multimarket contact theory, Coccorese and Pellecchia (2009) find that banks' profitability is positively related to the average number of contacts, and is higher for those credit institutions experiencing more links. The striking contrast between these two studies is most likely attributable to the dissimilar time periods under investigation and the differences between the models and their approach (Coccorese and Pellecchia, 2009, p. 263).

In what follows, we explore the relationship between multimarket contact and market performance in the Italian banking industry by estimating a Bresnahan-Lau simultaneous equation model of demand and supply for loans where the parameter measuring the degree of competition is built as a function of multimarket contact. For the purpose, we use panel data regarding 20 regional markets over 13 years. This framework allows to directly capture the possible influence of multimarket linkages among banks on the competitiveness of regional markets, and hence on the loan price.

To our knowledge, this is the first attempt to employ such methodology for the banking sector, and one of the first in the literature on market power. Jans and Rosenbaum (1996) develop an analogous Bresnahan-Lau model for the U.S. cement industry, showing that the greater the concentration in other markets where firms from a particular market meet, the higher the price in that particular market. However, they employ linear equations for both demand and marginal cost, while Perloff and Shen (2012) demonstrate that estimates based on the linear model inherently suffer from a severe multicollinearity problem that makes estimates of all the parameters in the optimality equation, including the market power parameter, unreliable, while this problem can be avoided if at least one of the equations has some other functional form. Gallet (1997) also estimates a Bresnahan-Lau two-equation model for the U.S. rayon industry in the 1930s, but the conduct parameter is supposed to be dependent on three variables aiming at capturing time variations only.

3. The model

The starting point of our analysis is the theoretical assumption that a profit-maximizing firm (here, a bank) chooses the level of output (here, loans) that equalizes marginal revenue to marginal cost. In case of a perfectly competitive market with *n* firms, marginal revenue corresponds to the demand price; in contrast, it is equal to the marginal revenue of the whole market if there is perfect collusion among the *n* firms.

Let us suppose the loan market demand, for a given region *r* at time *t*, is

$$Q_{rt} = Q_{rt}(P_{rt}, X_{rt}, \delta), \tag{1}$$

where Q_{rt} is the aggregate level of loans, P_{rt} is the interest rate on loans charged by local banks, X_{rt} is a vector of exogenous variables shifting the demand curve, and δ represents a vector of unknown parameters to be estimated.

The industry's *true* marginal revenue function corresponds to

$$MR_{rt} = P_{rt} + \frac{Q_{rt}(\cdot)}{\frac{\partial Q_{rt}(\cdot)}{\partial P_{rt}}}, \tag{2}$$

while the firm's *perceived* marginal revenue function for the generic bank *i* operating in that region, and supplying the quantity of loans q_{irt}, may be written as

$$MR_{irt} = P_{rt} + \lambda_{irt} \frac{q_{irt}(\cdot)}{\frac{\partial Q_{rt}(\cdot)}{\partial P_{rt}}}. \tag{3}$$

The positive unknown parameter λ_{irt} (to be estimated) measures the competitiveness of oligopoly conduct. It ranges from 0 to 1: when $\lambda_{irt} = 0$, each bank acts as though marginal revenue coincides with market demand, so that price equals marginal cost and the behaviour is perfectly competitive; when $\lambda_{irt} = 1$, banks choose price and output according to the industry marginal revenue curve, implying joint monopoly or perfect collusion. Intermediate values of λ_{irt} indicate various degrees of imperfect competition or market power, and aggregate output will be less than the competitive value. Should λ be negative, it could reflect an unsustainable deviation from long-run equilibrium characterized by pricing below marginal cost; hence, aggregate output will exceed the competitive norm (Gruben and McComb, 2003; Shaffer, 2004).

The overparametrization of this model can be solved by aggregation. Hence, we can use regional industry data for both demand and cost variables. In this framework, the industry's marginal cost function MC_{rt} is to be interpreted as the horizontal summation of the *n* banks' marginal cost functions (Shaffer, 2004, pp. 292-293). We are then left with a single parameter λ_{rt}, which measures the average conduct of the banks operating in market *r* at time *t* (Bresnahan, 1989, pp. 1016-1017).

After aggregating for the *n* banks in the market, the equilibrium condition $MR = MC$ can be written as

$$P_{rt} + \lambda_{rt} \frac{Q_{rt}(\cdot)}{\frac{\partial Q_{rt}(\cdot)}{\partial P_{rt}}} = MC_{rt}, \tag{4}$$

or else

$$P_{rt} - MC_{rt} = -\frac{\lambda_{rt}}{\eta_{rt}}, \tag{5}$$

where $\eta_{rt} = \frac{\partial Q_{rt}(\cdot)}{\partial P_{rt}} \frac{1}{Q_{rt}(\cdot)}$ is the semi-elasticity of market demand with respect to price.

Equation (5) provides a measure of the deviation of price from marginal cost *MC*, and therefore from the competitive price level. Shaffer (1993) demonstrates that $-\lambda$ is also a local (or linear) approximation of the percentage deviation of total industry output from the optimal level of production that characterizes the competitive equilibrium.

Since our aim is to investigate the source of market power, and particularly to check whether multimarket contact facilitates collusion among banks, our index of market power λ_{rt} is specified as a function of multimarket contact among banks operating in region *r* at time *t*, MMC_{rt}, and local market concentration, here measured by means of the Herfindhal-Hirschman index, HHI_{rt} (Jans and Rosenbaum, 1996; Fernandez and Marin, 1998):

$$\lambda_{rt} = \lambda_{rt}(MMC_{rt}, HHI_{rt}). \tag{6}$$

For each region, the Herfindhal-Hirschman index (HHI) ranges between 0 and 1 and has been calculated from the branch distribution of banks (for details, see next Section). It is included to test the possibility that, apart from multimarket links, concentration within regional markets allows local banks to raise price above marginal cost.

Among the advantages of this particular NEIO approach, it is worth to stress the fact that it provides an easily interpreted test statistic, and the possibility to use industry aggregate data (Shaffer, 2004, p. 294). Besides, the model does not rely on any particular definition of local banking markets within a country. In a sample spanning multiple markets (here, the Italian regions), the estimate of λ would represent the average degree of market power of the banks across those separate markets (Shaffer, 2001, p. 88). Finally, we do not expect any biased estimation of the market power parameter, since our sample spans a complete market rather than only a subset of the relevant industry (Shaffer, 2001, p. 90).

4. Multimarket contact measures

A peculiarity of banking industries is that one bank can meet another bank in many geographical markets because of the possibility of opening branch offices in several cities or towns. Hence, to build our multimarket contact measures, we follow a procedure similar to Evans and Kessides (1994) and Jans and Rosenbaum (1996).

Suppose that N banks operate in (some or all of) M regional markets. For a given year, let $BR = [br_{ij}]$ be the $N \times M$ matrix describing the geographical distribution of banks' branches, whose generic element br_{ij} (where $i = 1,\ldots,N$ indexes banks, and $j = 1,\ldots,M$ indexes markets) is the number of branches of bank i in market j.

Starting from matrix BR, we can build the $N \times M$ binary matrix $U = [u_{ij}]$ checking the presence of each bank in the various markets. In particular, it is $u_{ij} = 1$ if $br_{ij} > 0$ (i.e. bank i operates in market j by means of one or more branches), while it is $u_{ij} = 0$ if $br_{ij} = 0$.

Matrix U allows to construct the $N \times N$ symmetric matrix $A = UU'$, whose generic element is

$$a_{kl} = \sum_{j=1}^{M} u_{kj} u_{lj} \ . \qquad (7)$$

Each *off-diagonal* element a_{kl} of matrix A measures the number of markets (regions) in which banks k and l meet, while the generic *diagonal* element a_{kk} quantifies the number of markets (regions) where bank k operates.

Finally, if N_j is the number of banks operating in market *j*, the total number of possible pairings between these banks (i.e. the overall number of contacts among them in this market) is given by $N_j(N_j - 1)/2$.

Our first market-level measure of multimarket contact, $MMC1_j$, is computed as the average number of contacts in regions other than *j* ("non-home" regions) per contact in region *j* ("home" regions), as regards the banks operating in this region. Formally:

$$MMC1_j = \frac{\sum_{k=1}^{M-1}\sum_{l=k+1}^{M} a_{kl}u_{kj}u_{lj} - N_j(N_j - 1)/2}{N_j(N_j - 1)/2}. \qquad (8)$$

Here the numerator represents the number of times banks operating in region *j* meet each other in regions other than *j*.

The shortcoming of this multimarket contact measure is that it does not take into account the size of banks or the concentration prevailing in the markets, but only the number of contacts, which have the same weight for all banks and across markets. Actually, the higher the concentration is (i.e. there are few banks with considerable market shares), the bigger the losses due to the retaliation from competitors are. Besides, in highly concentrated markets banks can attain implicit collusion more easily and with less coordination costs. Hence, we need to develop some new index of multimarket contact that also captures the ability of banks to exploit it. For the purpose, in line with Jans and Rosenbaum (1996), we build two additional multimarket contact indicators.

The first alternative measure, denoted as $MMC2_j$, takes into account banks' market shares (measured in percentage values). Let $MS = [ms_{ij}]$ be the $N \times M$ matrix whose generic element is the market share of bank *i* in region *j*. Our matrix *MS* is built starting from the geographical distribution of branches (source: Bank of Italy). Therefore:

$$ms_{ij} = \frac{br_{ij}}{\sum_{i=1}^{N_j} br_{ij}} \cdot 100. \qquad (9)$$

For each market *j*, let $B^{(j)}$ denote the $N \times N$ symmetric matrix whose generic element $b_{kl}^{(j)}$ is given by the sum of market shares of banks *k* and *l* in all "non-home" markets (other than *j*) in which they meet:

$$b_{kl}^{(j)} = \sum_{p \neq j} \left(ms_{kp} + ms_{lp}\right) u_{kp}u_{lp}. \qquad (10)$$

Our $MMC2_j$ measure is then calculated as the average sum of market shares in markets other than j per contact in the same markets, as regards the banks operating in market j. Formally:

$$MMC2_j = \frac{\sum_{k=1}^{M-1}\sum_{l=k+1}^{M} b_{kl}^{(j)} u_{kj} u_{lj}}{\sum_{k=1}^{M-1}\sum_{l=k+1}^{M} a_{kl} u_{kj} u_{lj} - N_j (N_j - 1)/2} . \qquad (11)$$

The second alternative measure, called $MMC3_j$, considers the regional values of the Herfindahl-Hirschman index in place of banks' market shares (with HHI ranging between 0 and 1). For the purpose, for each market j we define the $N \times N$ symmetric matrix $C^{(j)}$ whose generic element is given by the concentration indexes characterizing the "non-home" markets (other than j) in which banks k and l meet:

$$c_{kl}^{(j)} = \sum_{p \neq j} HHI_p u_{kp} u_{lp} . \qquad (12)$$

As for market shares, the HHI of each region has been derived starting from banks' market shares ms_{ij} as calculated from the branch distribution:

$$HHI_j = \sum_{i=1}^{N_j} ms_{ij}^2 . \qquad (13)$$

With reference to the banks operating in market j, the $MMC3_j$ measure is given by the average HHI of "non-home" markets per contact in the same markets:

$$MMC3_j = \frac{\sum_{k=1}^{M-1}\sum_{l=k+1}^{M} c_{kl}^{(j)} u_{kj} u_{lj}}{\sum_{k=1}^{M-1}\sum_{l=k+1}^{M} a_{kl} u_{kj} u_{lj} - N_j (N_j - 1)/2} . \qquad (14)$$

By means of a numerical example, the Appendix shows how *MMC*1, *MMC*2 and *MMC*3 are calculated, while Figure 1 outlines the trend of the above measures (yearly averages) in the sample period. The upward trend of *MMC*1 (Figure 1, panel a) can be largely attributed to the fact that more widespread branch networks (originating from the growth of the number of offices per bank), joint with the various M&A's, have allowed banks to meet each other more frequently. On the contrary, the *MMC*2 measure (Figure 1, panel b) follows a downward path because the increasing presence of banks in the various regions has caused a generalized drop in their average market shares (and hence in each pairing's market share). Finally, *MMC*3 does not show a clear tendency (Figure 1, panel c): however, it appears correlated with the average regional HHI (Figure 1, panel d), whose trend shows a pattern of broad decline, again due to the branch growth. Lastly, in the various panels it is evident an abrupt change from 2007 to 2008 attributable to the merger between Banca Intesa and San Paolo IMI (which gave rise to Intesa Sanpaolo) as well as the buyout of Capitalia by Unicredit, both of which were effective in 2007.

INSERT FIGURE 1 ABOUT HERE

Figure 2 summarizes the regional values of the *MMC*'s and the HHI by quartiles (panels a to d). Multimarket contact measures are generally lower in the North and the Centre, especially the weighted ones (*MMC*2 and *MMC*3). The same happens for the HHI, meaning that a lower number of banks and/or local offices operates in the South compared to the other regions.

INSERT FIGURE 2 ABOUT HERE

5. Econometric specification

The estimation of the market power index λ requires a simultaneous model that comprises the demand equation (1) and the supply relation (5) including the λ_{rt} specification (6).

For region r and year t, the demand function takes the following semi-logarithmic form (Lopez et al., 2002; Captain et al., 2007; Coccorese, 2008):

$$\ln Q_{rt} = a_1 P_{rt} + a_2 Z_t + a_3 Y_{rt} + a_4 POP_{rt} + a_5 BR_{rt} + a_6 TIME + \alpha_r \; . \tag{15}$$

In (15), Q_{rt} is the amount of loans, and P_{rt} is the average market loan rate. Furthermore, Z_t is the interest rate of 3-month government bonds, a proxy for the price of a substitute for bank loans; Y_{rt} represents the regional Gross Domestic Product, which controls for the level of local aggregate demand; POP_{rt} measures the regional population, a variable that accounts for the absolute market size; BR_{rt} is the regional number of branches of the banking system, which helps to consider the banks' network size effect. Finally, *TIME* is included to capture possible trend effects, while α_r represents the constant associated to each region.

With respect to the double-logarithmic functional form, the semi-logarithmic demand function has the advantage of not imposing constant elasticities, which might not be a suitable assumption when dealing with time-series data. In addition, the estimated coefficient a_1 directly provides a measure of the average value of the price semi-elasticity of demand η.

In the supply relation (5), our assumption is that marginal cost originates from a translog cost function (a common hypothesis in the most recent banking studies) with two generic inputs (deposits, W_1, and labour, W_2) and one output (loans, q):

$$\ln C_{irt} = b_0 + b_Q \ln q_{irt} + \frac{b_{QQ}}{2}(\ln q_{irt})^2 + b_1 \ln W_{1rt} + b_2 \ln W_{2rt} + \ln q_{irt}\left(b_{Q1} \ln W_{1rt} + b_{Q2} \ln W_{2rt}\right) + \\ + \frac{1}{2}\left[b_{11}(\ln W_{1rt})^2 + b_{22}(\ln W_{2rt})^2 + b_{12} \ln W_{1rt} \ln W_{2rt} + b_{21} \ln W_{2rt} \ln W_{1rt}\right] + \\ + b_T \ln TIME + \frac{b_{TT}}{2}(\ln TIME)^2 + \ln TIME\left(b_{QT} \ln q_{irt} + b_{1T} \ln W_{1rt} + b_{2T} \ln W_{2rt}\right) , \qquad (16)$$

where C_{irt} and q_{irt} are bank i's total costs and output, respectively, W_{1rt} and W_{2rt} are the exogenous prices of the inputs, and $TIME$ is included to capture the possible effects of technological change over time (Zardkoohi and Fraser, 1998).

By the symmetry condition, it must be $b_{12} = b_{21}$. Furthermore, linear homogeneity in input prices requires that $b_1 + b_2 = 1$, $b_{11} + b_{12} = 0$, $b_{21} + b_{22} = 0$, $b_{Q1} + b_{Q2} = 0$, and $b_{1T} + b_{2T} = 0$. In order to impose these conditions, we divide total costs and factor prices by W_{2rt}. Finally, we aggregate for the banks operating in each of the regional markets.

All the above leads to the following regional translog cost function:

$$\ln\left(C_{rt} / W_{2rt}\right) = b_0 + b_Q \ln Q_{rt} + \frac{b_{QQ}}{2}(\ln Q_{rt})^2 + b_1 \ln\left(W_{1rt} / W_{2rt}\right) + \\ + b_{Q1} \ln Q_{rt} \ln\left(W_{1rt} / W_{2rt}\right) + \frac{b_{11}}{2} \ln\left(W_{1rt} / W_{2rt}\right)^2 + b_T \ln TIME + \frac{b_{TT}}{2}(\ln TIME)^2 + \\ + b_T \ln TIME + \frac{b_{TT}}{2}(\ln TIME)^2 + \ln TIME\left[b_{QT} \ln Q_{rt} + b_{1T} \ln\left(W_{1rt} / W_{2rt}\right)\right] . \qquad (17)$$

As a result, the marginal cost function is:

$$MC_{rt} = \frac{C_{rt}}{Q_{rt}}\left[b_Q + b_{QQ} \ln Q_{rt} + b_{Q1} \ln\left(W_{1rt} / W_{2rt}\right) + b_{QT} \ln TIME\right] . \qquad (18)$$

The price of funds, W_{1rt}, is measured through the regional interest rate on deposits, while the price of labour, W_{2rt}, is proxied by means of the average regional wage rate of the financial intermediation sector. Total cost, C_{rt}, is given by the sum of the interest expenses (computed as the product of the deposit rate and the amount of deposits) and the labour costs (assumed equal to the salaries paid by financial intermediaries), again at the regional level.

Substituting (18) in (5), rearranging and taking into consideration the semi-logarithmic nature of the demand equation, we can write:

$$P_{rt} = \frac{C_{rt}}{Q_{rt}}\left[b_Q + b_{QQ}\ln Q_{rt} + b_{Q1}\ln\left(W_{1rt}/W_{2rt}\right) + b_{QT}\ln TIME\right] - \frac{\lambda_{rt}}{a_1}\,. \tag{19}$$

Multimarket contact and market concentration in region *r* at time *t* are supposed to affect the index of market power λ_{rt} as follows:

$$\lambda_{rt} = \lambda_0 + \lambda_1 MMC_{rt} + \lambda_2 HHI_{rt} + \lambda_3 MMC_{rt} \times HHI_{rt}\,. \tag{20}$$

While the inclusion of *MMC* allows to test the main hypothesis of our paper, the regional value of *HHI* is added to control for the possibility that banks in more concentrated markets exploit a greater market power, as the SCP paradigm maintains. The interaction term *MMC*×*HHI* helps to establish whether the effect of multimarket contact (local concentration) on the degree of collusion is more or less sizeable as local concentration (multimarket contact) increases.

Substituting (20) in (19), we get:

$$P_{rt} = \frac{C_{rt}}{Q_{rt}}\left[b_Q + b_{QQ}\ln Q_{rt} + b_{Q1}\ln\left(W_{1rt}/W_{2rt}\right) + b_{QT}\ln TIME\right] - \\ - \frac{\left(\lambda_0 + \lambda_1 MMC_{rt} + \lambda_2 HHI_{rt} + \lambda_3 MMC_{rt} \times HHI_{rt}\right)}{a_1}\,. \tag{21}$$

In conclusion, the system we are going to estimate is formed by equations (15) and (21). According to Lau (1982), a necessary and sufficient condition for the correct identification of the parameter indexing the degree of market power (here, λ_{rt}) in a system of demand and cost equations is that the demand function must *not* be separable in at least one exogenous variable that is included in the demand function but excluded from the marginal cost function. This condition is met by our semi-logarithmic demand function, since, for example, it is $\frac{\partial^2 Q_{rt}}{\partial P_{rt}\partial POP_{rt}} = a_1 a_4 Q_{rt} \neq 0$.

6. Empirical evidence and discussion

Our dataset covers the twenty Italian regions[6] for the years 1997-2009. Regional data have been collected from the Central Bank of Italy and the Italian Statistical Institute (Istat). Economic figures have been converted into real values by means of the regional Gross Domestic Product deflator, with 2000 as the base year. Table 1 shows some descriptive statistics of the data.

INSERT TABLE 1 ABOUT HERE

In order to get precise and efficient estimates, we employ nonlinear three-stage least squares. Conforming to the standard practice, we use all exogenous variables as instruments (including time trend and regional dummies). Due to the endogeneity of Q_{rt} and P_{rt}, we also use their first lagged values as instruments, so as to deal with possible problems of correlation between these variables and the error terms. Additional instruments are the number of employees, and the values of regional deposits, consumption and investment; all of them proxy for (different aspects of) market size.

With reference to the behavioural parameter λ_{rt}, we estimate different specifications of our model. First, we treat it as a constant, depending on neither multimarket contact nor market concentration (i.e. $\lambda_{rt} = \lambda_0$), in line with the customary Bresnahan-Lau mark-up test. Next, for each of our three measures of multimarket linkages we estimate the system progressively adding the explanatory variables of λ_{rt} (*MMC*, *HHI* and *MMC*×*HHI*). This procedure allows us to analyse how the coefficients of interest change, and therefore gives an idea about the sensitivity of λ_{rt} to different model specifications. The results of all estimations are presented in Table 2.

INSERT TABLE 2 ABOUT HERE

It is self-evident that our results are robust to the use of different models. The coefficients of the explanatory variables in the demand equation are fairly similar in all specifications, and four of them are always significant at least at the 5% level. The overall fit is quite good, as the adjusted R^2 values are very high. The price of loans, *P*, exhibits a negative sign, which confirms the downward slope of the demand curve. The coefficient of *Z* is positive, showing that it is a good measure of the price of a substitute of bank loans. In absolute value, the calculated elasticity of *P*, $\varepsilon_{Q,P}$, is always greater than the elasticity of *Z*, $\varepsilon_{Q,Z}$ (see Table 2): this result provides evidence that, being the cross-price elasticity smaller than the own-price elasticity, banks are able to soften price competition

[6] According to the classification of the Central Bank of Italy, they can be divided into four macro-areas: North-West (Piemonte, Valle D'Aosta, Lombardia, Liguria); North-East (Veneto, Trentino Alto Adige, Friuli Venezia Giulia, Emilia Romagna); Center (Toscana, Umbria, Marche, Lazio); South and Islands (Abruzzi, Molise, Campania, Basilicata, Puglia, Calabria, Sicilia, Sardegna).

thanks to product or service differentiation and the provision of fringe services. In addition, since the value of the own-price elasticity of demand ranges between -0.21 and -0.27, equilibrium is set in the inelastic portion of the demand curve, This is in line with other studies analyzing the banking industry (e.g.: Shaffer and DiSalvo, 1994; Shaffer, 2004; Coccorese, 2005, 2009; Richards et al., 2008).

The coefficient associated to *POP* is positive and always highly significant, indicating that wider markets guarantee banks a higher loan demand. The variables *Y* and *BR* exhibit positive but non-significant coefficients: hence, during the years under consideration GDP and banks' branch network did not play a major role in stimulating loan demand. Finally, the positive and significant value of the time trend coefficient suggests that loan size has increased in the sample period.

Turning to the supply equation, we first observe that it fits the data fairly well, considering that the estimated regression equation explains between 71% and 77% of the variability of *P*. The coefficients of all variables of the marginal cost function (except $\ln Q$) are generally significant. Their estimated values allow us to establish that on average the Italian banks exhibit economies of scale: actually, at the sample means the calculated average total cost per euro of loan (0.0537 euro) is always higher than the corresponding estimated marginal cost (ranging between 0.0180 and 0.0329 euro).[7] Again, this finding is consistent with the results of previous studies on cost functions of Italian banks (e.g.: Conigliani et al., 1991; Altunbas and Molyneux, 1996; Simper, 1999; Girardone et al., 2004).

In the same equation we find the parameters that allow to measure the market power of Italian banks. In the specification where multimarket contact is not considered, and hence λ is treated as constant throughout the whole country (first column), the market power parameter is equal to 0.1465, meaning that during the sample period banks' perceived marginal revenue has been only about 15 per cent of the marginal revenue that would be taken into consideration by a monopolistic firm or a cartel. Having tested that λ is significantly different from zero and one, we can reject the hypotheses of both perfect collusion and perfect competition. The value $\lambda = 0.1465$ is consistent with the estimation provided by Coccorese (2008) for the Italian banking sector in the period 1995-2004 (0.1497), and supports the view that competition among Italian banks is nonetheless intense in spite of the consolidation trend that has characterized this industry in the last twenty years. Considering that in a Cournot equilibrium it is $\lambda = 1/n$ (Bresnahan, 1982, p. 88), our point estimate of the conduct parameter also corresponds to a symmetric Cournot oligopoly with about seven identical firms whose market shares are just equal to λ.

[7] We are aware that the above comparison would require the estimated, rather than the calculated, average total cost, but this figure is not available here, because our estimation of the marginal cost – Equation (18) – provides some but not all the parameters of the translog total cost function, as formulated in Equation (17).

Figure 3 depicts the average situation of Italian regional markets when $\lambda = 0.1465$. The various curves have been built using the estimation results of the first column of Table 2. Point E indicates the equilibrium that has characterized the Italian banking industry in the sample period, where marginal cost equals perceived marginal revenue. We note that the corresponding calculated value of Q (loans) is 19995.5 million euro, very close to the median value of our sample (21232.7 million euro). We also observe that banks did not behave as profit-maximizing firms, because they fixed the equilibrium quantity where the industry's marginal revenue is negative, implying that the market demand they faced was perceived to be inelastic. This outcome is in line with our estimation of the own-price elasticity of demand.

INSERT FIGURE 3 ABOUT HERE

Profit maximization would have required to set the quantity that equates marginal cost and the industry's marginal revenue, i.e. point M (where $Q = 8497.1$ million euro). Since this first-order condition for static profit maximization is not satisfied, we can reject the hypothesis of a monopoly conduct of Italian banks. Quite to contrary, their behaviour has been much closer to that of a firm operating in a perfectly competitive environment, here corresponding to point C (where the calculated Q is 23159.8 million euro), where firms do not perceive a difference between their marginal revenue functions and demand function and act as price takers by setting the price equal to marginal cost. Under this respect, Figure 3 makes clear that λ is a measure of the distance between the perceived marginal revenue curve and the demand curve relative to the distance between industry's marginal revenue curve and the demand curve. Besides, the deviation of the calculated actual output from the calculated competitive output amounts to -13.7%, a value very close to $-\lambda$ expressed as a percentage (see Section 2).

We have now to evaluate the influence of multimarket contact on the degree of competition in the Italian banking market. Because of the specification of λ in Equation (20), we are able to estimate different λ_{rt}'s for each region and year. The last rows of Table 2 report the average, the minimum and the maximum regional value of λ, as calculated for all models and multimarket contact measures. None of the indexes is negative, and all of them fall into the theoretical boundaries of zero and one. Furthermore, they do not differ much across model specifications.

Let us first analyze the results obtained using the first measure of multimarket contact, *MMC*1 (second to fourth columns of Table 2). Adding *MMC*1 and *HHI* in the market power equation without their interaction (second and third column) does not change the evidence presented so far: their coefficients, λ_1 and λ_2, are not significant, and the values of the different λ_{rt}'s are essentially driven by the constant term λ_0. Actually, as the bottom lines of Table 2 make clear, the average estimated market power index is quite similar to the value we have got using the standard approach (i.e. with a unique λ), and the range of variation between regions and years is rather small. On the other hand, the inclusion of the interaction between multimarket contact and regional market concentration, $MMC1 \times HHI$, makes all variables significant at the 5% level (fourth column). Now the coefficients of $MMC1$ and HHI exhibit a positive sign, while the coefficient of $MMC1 \times HHI$, λ_3, is negative.

We deduce that multimarket contact helps to mitigate rivalry among banks, because it contributes to increase the market power index. This evidence corroborates the view according to which firms that meet in more markets increase their power to collude, and is in line with the results of Coccorese and Pellecchia (2009), although the latter have been obtained employing a different theoretical framework. Also, market power is directly linked with local market concentration: conforming to the SCP paradigm, banks in more concentrated markets enjoy greater market power.

However, the significance of the above two variables is associated to the insertion of their interaction. To our purpose, this means that multimarket contact has a role in increasing market power only when it is joint with home market concentration; in other words, what matters is not the mere number of contacts, but their importance with reference to each market. Since $\lambda_1 > 0$ and $\lambda_3 < 0$, banks' market power is higher if multimarket linkages increase, but a given value of contacts has less and less effect on the degree of market power as local concentration becomes higher.

One explanation for this outcome is that, when regional concentration is low, coordination among banks is difficult, so tacit collusion comes in handy and multimarket linkages constitute an effective coordination device that enables banks to raise prices. Their relevance however lessens as the ease of collusion, as measured by HHI, increases. Along with Jans and Rosenbaum (1996), we can also say that $MMC1$ captures the diversification dimension of multimarket contact, but not necessarily the ability of banks to use it to increase their market power, which needs also favourable structural conditions. All in all, multimarket contact and local market concentration tend to act as substitutes in guaranteeing a higher degree of market power to banks. Such finding is similar to what has been observed by Fernandez and Marin (1998) for the Spanish hotel industry, and by Coronado et al. (2008) for the pharmaceutical markets of nine OECD countries, but contrasts with the evidence of Jans and Rosenbaum (1996) concerning the U.S. cement industry, where the interaction between multimarket linkages and market concentration exhibits a positive sign.

Since in our model $\partial\lambda_{rt}/\partial MMC1 = \lambda_1 + \lambda_3 HHI_{rt}$, we can calculate how λ_{rt} changes as HHI varies. The Herfindahl-Hirschman index ranges between 0.0302 and 0.3697 (see Table 1), so the predicted impact of multimarket contact on the market power index goes from 0.0053 to -

0.0214, remaining positive until HHI is equal to 0.0976 (i.e. 0.00766/0.07847). Since this value corresponds approximately to the third quartile of our sample, we conclude that in many Italian regions a higher market concentration generally tends to increase banks' market power via multimarket linkages.

A look at the estimated values of λ_{rt}, as calculated starting from the results of the fourth column of Table 2, makes clear that banking markets of Southern regions are characterized by a lower degree of competition (see Table 3): particularly, the highest average values of the market power index for the sample period are found in Calabria, Sicilia, Molise, Basilicata and Campania, while the contrary happens in Northern regions (the lowest values correspond to Trentino Alto Adige, Lombardia, Emilia, Veneto and Toscana). As Coccorese (2008) also notes, the evidence is that competition among banks appears to be stronger (i.e. λ is lower) in those regions characterized by a more dynamic economy, in terms of lower unemployment rates, lower bad loans to total loans ratio, and greater per capita GDP. Besides, the average level of the yearly λ shows a generalized decline up to 2007, and a sharp increase in 2008 and 2009, corresponding to the period of the world financial crisis.

INSERT TABLE 3 ABOUT HERE

However, we also discover a sort of convergence of the various λ_{rt}'s: actually, regions with lower average λ's are generally characterized by a growth of this index in the period under consideration, while there is a widespread reduction of λ in those regions where its average level is higher. In order to assess whether banks' market power in each region is really converging to the national average level, we have used the sigma-convergence test (Barro and Sala-I-Martin, 1992, 1995) in a specification suitable for panel data (Parikh and Shibata, 2004; Weill, 2011):

$$\Delta W_{r,t} = \delta_1 W_{r,t-1} + \gamma_r \,, \qquad (22)$$

with $\Delta W_{r,t} = W_{r,t} - W_{r,t\text{-}1}$ and $W_{r,t} = \ln\lambda_{r,t} - \text{mean}(\ln\lambda_t)$, while γ_r is a vector of dummies that is added to capture the regional effects. There is a sigma-convergence if $\delta_1 < 0$. Since in our estimation of (22) we get $\delta_1 = -0.3394$ (significantly different from zero at the 1% level), we conclude that in the sample period there has been a convergence in the various regional banking market power indexes, because the dispersion of the mean values of λ between regions has reduced.[8]

[8] The estimation results are available from the authors upon request.

We now turn to consider how our second index of multimarket contact, *MMC*2, which measures the average market share in non-home regions of each pair of banks operating in region *r*, impacts over the market power index. As Table 2 shows (fifth to seventh columns), the most striking evidence is that the coefficients of both *HHI* and *MMC*2×*HHI* are never significant, while *MMC*2 exhibits a positive a significant coefficient (at the 5% level in the fifth and sixth columns, at the 10% level in the seventh column). This result states that multimarket linkages play an important role in determining the degree of competition or cooperation among banks as long as they are weighted with their relevance within non-home markets. On the other hand, when using *MMC*1, based on the simple number of contacts, we have discovered a significant (positive) influence of multimarket contact on the market power index only when we interact this variable with home market concentration, i.e. when we explicitly take into account a variable measuring local market structure. Summing up, banks' market power becomes stronger with multimarket contact, and is also as much higher both as firms in their non-home regions are larger and as home regions are more concentrated.

The last three estimations of our model, which use *MMC*3 as multimarket contact variable (based on the average HHI of non-home regions), further corroborate the above conjecture (see eighth to tenth column of Table 2), because the only significant coefficient is again λ_1, whose sign is positive. Hence, the more concentrated are non-home regions where the typical pair of banks meets, the stronger is their market power in the home region, where they are therefore able to set a higher price.

Thus, there is a robust evidence that concentration in both home and non-home regions, combined with multimarket linkages, can ease market collusion.

To assess the effect of multimarket contact on the market power index, and hence on price-cost margins, we have calculated the price elasticity with respect to our MMC measures, $\varepsilon_{P,MMC}$, for each estimated model (see Table 2). It is always statistically indistinguishable from zero when using *MMC*1, while is positive and significant (generally at the 5% level) in the estimations employing either *MMC*2 and *MMC*3. With *MMC*2, the elasticity is about 0.099, meaning that a 10% increase in the sum of the market shares of the average pair of banks causes an increase of about 1% in the value of the interest rate charged to customers by banks. So, for example, if *MMC*2 passed from the median value (3.8162) to the third quartile (4.9842), corresponding to +31%, the price would increase by about 3%. Similarly, the estimated elasticity of *P* with respect to *MMC*3 varies between 0.75 and 0.80, so an increase of *MMC*3 from the median (0.0750) to the third quartile (0.0793), amounting to +5.7%, would generate a price growth of 4.3%-4.6%. Given that the third quartile of

P (8.09) is 13.2% higher than its median (7.15), we deduce that the effect of MMC on loan price – although not high – can not be considered negligible.

This conclusion gets along with the findings of Coccorese and Pellecchia (2009), whose results show that in the Italian banking industry mutual forbearance allows higher profits, thus having a detrimental effect on competition, and provide support for the hypothesis that multimarket contact is able to reduce the level of market competition. Furthermore, as our estimated coefficients of the multimarket contact variables are generally significant, we conclude that omitting this variable can lead to a misspecification of the model.

7. Conclusion

As the standard economic literature maintains, the fact that firms meet repeatedly in different markets may induce them to promote and sustain tacit collusion in each single market, because the prospect of gains from deviation in a local market is not worth the risk of warfare in all the other markets. This means that multimarket contact can adversely affect the intensity of competition in the considered industry.

In this paper we have empirically tested the theory of mutual forbearance for the Italian banking industry by assessing the effect of multimarket contact on the degree of market power of banks, and hence on pricing. This has been done by means of a simultaneous equation model, in which the divergence of price from marginal cost has been considered as a function of multimarket linkages (calculated in three alternative ways). The model has been estimated using data from the twenty Italian regions for the years 1997-2009.

According to our results, the estimated market power parameters generally confirm the evidence of other previous studies according to which competition among Italian banks is fairly intense notwithstanding the consolidation trend that has characterized this industry in the last years.

Multimarket links are found to be positively and significantly correlated to the market power index, and hence to banks' pricing. The effect of multimarket contact is also connected to market concentration. Particularly, if multimarket contact is measured as the average number of contacts in regions different from the one where a pair of banks meets, market power and pricing are directly linked to multimarket linkages and local market concentration, but a given number of average linkages exerts less and less effect on market power as regional concentration becomes higher. Our explanation is that multimarket contact and local market concentration are substitutes in guaranteeing a higher degree of market power to banks.

In addition, when we use indexes of multimarket contact that are weighted by either the average market share in non-home regions of each pair of banks and the average HHI of non-home regions, we find evidence that the combination of multimarket linkages and non-home market concentration can facilitate market collusion. These results are robust to changes in model specification. In general, the calculated effect of multimarket contact on loan price is not minor.

From a policy perspective, our results indicate that the latest trend towards consolidation in the banking sector might reduce competition among banks especially because of the increased geographic overlap of banks' branch networks, which has brought about more frequent multimarket linkages among credit institutions. Hence, multimarket contact should be taken into consideration by antitrust authorities, given that it is able to increase banks' market power even when market concentration remains unchanged. Should the rise in the number of contacts among banks not be outweighed by the development of pro-competitive devices (e.g. adopting technological innovations in the provision of services to clientele on a larger scale, favouring the mobility of customers, enhancing the entry of new competitors, etc.), the more concentrated structure of banking markets could produce less advantageous conditions (with welfare losses) in a vital industry for the economic performance and growth of all countries.

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APPENDIX – *Calculation of the multimarket contact measures: an example*

Hereafter we provide a computational example of how our multimarket contact measures, as defined in Section 4, have been determined. Let us suppose that, in a given year, 5 banks operate in 4 markets, and that the geographical distribution of their branches is as follows:

$$BR = \begin{bmatrix} 3 & 9 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 4 & 0 & 3 \\ 6 & 0 & 9 & 5 \\ 1 & 2 & 1 & 8 \end{bmatrix}$$

As an example, $br_{12} = 9$ means that bank 1 has 9 branches in market 2.

The binary matrix U is therefore:

$$U = \begin{bmatrix} 1 & 1 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 1 & 0 & 1 \\ 1 & 0 & 1 & 1 \\ 1 & 1 & 1 & 1 \end{bmatrix}$$

The i-th row indicates the markets where bank i operates, while the j-th column specifies which banks are located in market j. Moreover, the sum of each column delivers the number of banks that are active in that particular market.

The symmetric matrix A comes out to be:

$$A = UU' = \begin{bmatrix} 2 & 1 & 1 & 1 & 2 \\ 1 & 1 & 1 & 0 & 1 \\ 1 & 1 & 2 & 1 & 2 \\ 1 & 0 & 1 & 3 & 3 \\ 2 & 1 & 2 & 3 & 4 \end{bmatrix}$$

In this matrix, $a_{11} = 2$ indicates that bank 1 owns branches in two markets (see the first row of matrix BR), while $a_{12} = 1$ means that bank 1 meets bank 2 in just one market (i.e. the second one, according to the second column of matrix BR).

For each market j, the computation of $MMC1_j$ needs both the upper (or the lower) triangular part of matrix A and the columns of matrix U. According to (8), we can write:

$$MMC1_1 = \frac{[(1\cdot1\cdot0+1\cdot1\cdot0+1\cdot1\cdot1+2\cdot1\cdot1+1\cdot0\cdot0+0\cdot0\cdot1+1\cdot0\cdot1+1\cdot0\cdot1+2\cdot0\cdot1+3\cdot1\cdot1)-3\cdot2/2]}{3\cdot2/2} = \frac{3}{3} = 1$$
$$MMC1_2 = \frac{[(1\cdot1\cdot1+1\cdot1\cdot1+1\cdot1\cdot0+2\cdot1\cdot1+1\cdot1\cdot1+0\cdot1\cdot0+1\cdot1\cdot1+1\cdot1\cdot0+2\cdot1\cdot1+3\cdot0\cdot1)-4\cdot3/2]}{4\cdot3/2} = \frac{2}{6} = 0.33$$
$$MMC1_3 = \frac{[(1\cdot0\cdot0+1\cdot0\cdot0+1\cdot0\cdot1+2\cdot0\cdot1+1\cdot0\cdot0+0\cdot0\cdot1+1\cdot0\cdot1+1\cdot0\cdot1+2\cdot0\cdot1+3\cdot1\cdot1)-2\cdot1/2]}{2\cdot1/2} = \frac{2}{1} = 2$$
$$MMC1_4 = \frac{[(1\cdot0\cdot0+1\cdot0\cdot1+1\cdot0\cdot1+2\cdot0\cdot1+1\cdot0\cdot1+0\cdot0\cdot1+1\cdot0\cdot1+1\cdot1\cdot1+2\cdot1\cdot1+3\cdot1\cdot1)-3\cdot2/2]}{3\cdot2/2} = \frac{3}{3} = 1$$

For example, $MMC1_3 = 2$ means that the banks operating in the third market (i.e. the fourth and fifth banks) meet twice in all the other markets (particularly, in the first and the fourth markets).

In order to calculate the various $MMC2_j$'s, we first need to build matrix MS containing banks' market shares ms_{ij} (in percentage values):

$$MS = \begin{bmatrix} 30 & 56.25 & 0 & 0 \\ 0 & 6.25 & 0 & 0 \\ 0 & 25 & 0 & 18.75 \\ 60 & 0 & 90 & 31.25 \\ 10 & 12.5 & 10 & 50 \end{bmatrix}$$

By means of (10), we then construct the various symmetric matrices $B^{(j)}$:

$$B^{(1)} = \begin{bmatrix} 0 & 62.5 & 81.25 & 0 & 68.75 \\ 62.5 & 0 & 31.25 & 0 & 18.75 \\ 81.25 & 31.25 & 0 & 50 & 106.25 \\ 0 & 0 & 50 & 0 & 181.25 \\ 68.75 & 18.75 & 106.25 & 181.25 & 0 \end{bmatrix} \quad B^{(2)} = \begin{bmatrix} 0 & 0 & 0 & 90 & 40 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 50 & 68.75 \\ 90 & 0 & 50 & 0 & 251.25 \\ 40 & 0 & 68.75 & 251.25 & 0 \end{bmatrix}$$

$$B^{(3)} = \begin{bmatrix} 0 & 62.5 & 81.25 & 90 & 108.75 \\ 62.5 & 0 & 31.25 & 0 & 18.75 \\ 81.25 & 31.25 & 0 & 50 & 106.25 \\ 90 & 0 & 50 & 0 & 151.25 \\ 108.75 & 18.75 & 106.25 & 151.25 & 0 \end{bmatrix} \quad B^{(4)} = \begin{bmatrix} 0 & 62.5 & 81.25 & 90 & 108.75 \\ 62.5 & 0 & 31.25 & 0 & 18.75 \\ 81.25 & 31.25 & 0 & 0 & 37.5 \\ 90 & 0 & 0 & 0 & 170 \\ 108.75 & 18.75 & 37.5 & 170 & 0 \end{bmatrix}$$

For instance, the element $b_{45}^{(1)}$ of matrix $B^{(1)}$ has been calculated in the following way:

$$b_{45}^{(1)} = (0+12.5)\cdot0\cdot1+(90+10)\cdot1\cdot1+(31.25+50)\cdot1\cdot1 = 181.25$$

Using either the upper or the lower triangular part of the $B^{(j)}$ matrices, joint with the columns of matrix U, we use (11) and finally get:[9]

[9] Note that the denominator of (11) is equal to the numerator of (8).

$$MMC2_1 = \frac{62.5\cdot 1\cdot 0 + 81.25\cdot 1\cdot 0 + 0\cdot 1\cdot 1 + 68.75\cdot 1\cdot 1 + 31.25\cdot 0\cdot 0 + 0\cdot 0\cdot 1 + 18.75\cdot 0\cdot 1 + 50\cdot 0\cdot 1 + 106.25\cdot 0\cdot 1 + 181.25\cdot 1\cdot 1}{3} = 83.33$$

$$MMC2_2 = \frac{0\cdot 1\cdot 1 + 0\cdot 1\cdot 1 + 90\cdot 1\cdot 0 + 40\cdot 1\cdot 1 + 0\cdot 1\cdot 1 + 0\cdot 1\cdot 0 + 0\cdot 1\cdot 1 + 50\cdot 1\cdot 0 + 68.75\cdot 1\cdot 1 + 251.25\cdot 0\cdot 1}{2} = 54.38$$

$$MMC2_3 = \frac{62.5\cdot 0\cdot 0 + 81.25\cdot 0\cdot 0 + 90\cdot 0\cdot 1 + 108.75\cdot 0\cdot 1 + 31.25\cdot 0\cdot 0 + 0\cdot 0\cdot 1 + 18.75\cdot 0\cdot 1 + 50\cdot 0\cdot 1 + 106.25\cdot 0\cdot 1 + 151.25\cdot 1\cdot 1}{2} = 75.63$$

$$MMC2_4 = \frac{62.5\cdot 0\cdot 0 + 81.25\cdot 0\cdot 1 + 90\cdot 0\cdot 1 + 108.75\cdot 0\cdot 1 + 31.25\cdot 0\cdot 1 + 0\cdot 0\cdot 1 + 18.75\cdot 0\cdot 1 + 0\cdot 1\cdot 1 + 37.5\cdot 1\cdot 1 + 170\cdot 1\cdot 1}{3} = 69.17$$

Here, $MMC2_3 = 75.63$ means that, in all the other markets different from the third one where the banks operating in the third market meet (i.e. the first and the fourth market), the average sum of market shares of these banks (i.e. the fourth and fifth banks) is equal to 75.63%.

The way of computing $MMC3_j$ is similar to that of $MMC2_j$, except that it is based on the Herfindahl-Hirschman indexes, which here are:

$$HHI = \begin{bmatrix} 0.46 & 0.3984 & 0.82 & 0.3828 \end{bmatrix}$$

From (12) we derive the matrices $C^{(j)}$:

$$C^{(1)} = \begin{bmatrix} 0 & 0.3984 & 0.3984 & 0 & 0.3984 \\ 0.3984 & 0 & 0.3984 & 0 & 0.3984 \\ 0.3984 & 0.3984 & 0 & 0.3828 & 0.7813 \\ 0 & 0 & 0.3828 & 0 & 1.2028 \\ 0.3984 & 0.3984 & 0.7813 & 1.2028 & 0 \end{bmatrix} \quad C^{(2)} = \begin{bmatrix} 0 & 0 & 0 & 0.46 & 0.46 \\ 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0.3828 & 0.3828 \\ 0.46 & 0 & 0.3828 & 0 & 1.6628 \\ 0.46 & 0 & 0.3828 & 1.6628 & 0 \end{bmatrix}$$

$$C^{(3)} = \begin{bmatrix} 0 & 0.3984 & 0.3984 & 0.46 & 0.8584 \\ 0.3984 & 0 & 0.3984 & 0 & 0.3984 \\ 0.3984 & 0.3984 & 0 & 0.3828 & 0.7813 \\ 0.46 & 0 & 0.3828 & 0 & 0.8428 \\ 0.8584 & 0.3984 & 0.7813 & 0.8428 & 0 \end{bmatrix} \quad C^{(4)} = \begin{bmatrix} 0 & 0.3984 & 0.3984 & 0.46 & 0.8584 \\ 0.3984 & 0 & 0.3984 & 0 & 0.3984 \\ 0.3984 & 0.3984 & 0 & 0 & 0.3984 \\ 0.46 & 0 & 0 & 0 & 1.28 \\ 0.8584 & 0.3984 & 0.3984 & 1.28 & 0 \end{bmatrix}$$

while the multimarket contact measures $MMC3_j$ are calculated as follows:

$$MMC3_1 = \frac{0.3984\cdot 1\cdot 0 + 0.3984\cdot 1\cdot 0 + 0\cdot 1\cdot 1 + 0.3984\cdot 1\cdot 1 + 0.3984\cdot 0\cdot 0 + 0\cdot 0\cdot 1 + 0.3984\cdot 0\cdot 1 + 0.3828\cdot 0\cdot 1 + 0.7813\cdot 0\cdot 1 + 1.2028\cdot 1\cdot 1}{3} = 0.5338$$

$$MMC3_2 = \frac{0\cdot 1\cdot 1 + 0\cdot 1\cdot 1 + 0.46\cdot 1\cdot 0 + 0.46\cdot 1\cdot 1 + 0\cdot 1\cdot 1 + 0\cdot 1\cdot 0 + 0\cdot 1\cdot 1 + 0.3828\cdot 1\cdot 0 + 0.3828\cdot 1\cdot 1 + 1.6628\cdot 0\cdot 1}{2} = 0.4214$$

$$MMC3_3 = \frac{0.3984\cdot 0\cdot 0 + 0.3984\cdot 0\cdot 0 + 0.46\cdot 0\cdot 1 + 0.8584\cdot 0\cdot 1 + 0.3984\cdot 0\cdot 0 + 0\cdot 0\cdot 1 + 0.3984\cdot 0\cdot 1 + 0.3828\cdot 0\cdot 1 + 0.7813\cdot 0\cdot 1 + 0.8428\cdot 1\cdot 1}{2} = 0.4214$$

$$MMC3_4 = \frac{0.3984\cdot 0\cdot 0 + 0.3984\cdot 0\cdot 1 + 0.46\cdot 0\cdot 1 + 0.8584\cdot 0\cdot 1 + 0.3984\cdot 0\cdot 1 + 0\cdot 0\cdot 1 + 0.3984\cdot 0\cdot 1 + 0\cdot 1\cdot 1 + 0.3984\cdot 1\cdot 1 + 1.28\cdot 1\cdot 1}{3} = 0.5595$$

Here, $MMC3_3 = 0.4212$ means that on average, for the banks operating in the third market, the concentration index HHI that characterizes all the other markets different from the third one where they meet amounts to 0.4212.

TABLE 1 – *Sample descriptive statistics (1997-2009)*

Variable	Description	Mean	Std. Dev.	Minimum	Maximum	Median
Q	Total regional loans (*million euro at 2000 constant prices*)	46281.0	72392.1	779.3	475825.4	21232.7
P	Regional loan rate (*percentage*)	7.45	1.61	4.66	13.66	7.15
Z	Interest rate of 3-month government bonds (*percentage*)	3.32	1.38	0.97	6.40	3.15
Y	Regional Gross Domestic Product (*thousand million euro at 2000 constant prices*)	60.55	58.26	3.18	268.57	33.09
POP	Regional population (*million units*)	2.90	2.31	0.12	9.78	1.84
BR	Regional number of banks' branches (*thousand units*)	1.49	1.34	0.08	6.71	0.96
TC	Total regional banks' costs for deposits and labour (*million euro at 2000 constant prices*)	1914.7	2237.3	49.68	12443.0	1062.9
ω_1	Regional interest rate on deposits (*percentage*)	1.63	0.88	0.54	4.49	1.45
ω_2	Average regional financial sector wage (*thousand euro at 2000 constant prices*)	50.28	3.55	42.27	62.31	50.14
*MMC*1	Multimarket contact measure #1 (*based on the number of contacts*)	2.7636	2.1249	0.0819	9.3810	2.2750
*MMC*2	Multimarket contact measure #2 (*based on the sum of market shares*)	4.1072	1.3777	2.0903	8.3891	3.8162
*MMC*3	Multimarket contact measure #3 (*based on the Herfindahl-Hirschman index*)	0.0756	0.0059	0.0649	0.0931	0.0750
HHI	Regional Herfindahl-Hirschman index (*calculated on banks' branches*)	0.0920	0.0621	0.0302	0.3697	0.0740

Number of observations: 260

TABLE 2 – *System estimation results*

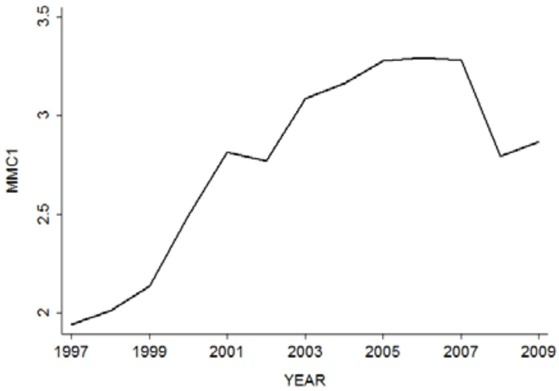
Variable		Without MMC	MMC1			MMC2			MMC3		
P	a_1	-0.03521 ***	-0.03524 ***	-0.03462 ***	-0.02547 **	-0.03561 ***	-0.03514 ***	-0.03506 ***	-0.02966 ***	-0.02817 ***	-0.02872 ***
		(-3.38)	*(-3.39)*	*(-3.32)*	*(-2.42)*	*(-3.41)*	*(-3.37)*	*(-3.37)*	*(-2.81)*	*(-2.67)*	*(-2.72)*
Z	a_2	0.03264 ***	0.03263 ***	0.03236 ***	0.02934 ***	0.03307 ***	0.03287 ***	0.03280 ***	0.02933 ***	0.02885 ***	0.02896 ***
		(3.24)	*(3.25)*	*(3.22)*	*(2.88)*	*(3.28)*	*(3.26)*	*(3.26)*	*(2.89)*	*(2.83)*	*(2.85)*
Y	a_3	0.00001	0.00003	0.00007	0.00056	-0.00026	-0.00022	-0.00022	0.00007	0.00017	0.00009
		(0.00)	*(0.01)*	*(0.02)*	*(0.20)*	*(-0.09)*	*(-0.08)*	*(-0.08)*	*(0.03)*	*(0.06)*	*(0.03)*
POP	a_4	0.47470 ***	0.47504 ***	0.47470 ***	0.46093 ***	0.47723 ***	0.47619 ***	0.47507 ***	0.47614 ***	0.47285 ***	0.47507 ***
		(4.43)	*(4.43)*	*(4.42)*	*(4.20)*	*(4.44)*	*(4.42)*	*(4.42)*	*(4.38)*	*(4.33)*	*(4.36)*
BR	a_5	0.02122	0.02127	0.02116	0.01871	0.02009	0.02001	0.02009	0.01920	0.01886	0.01918
		(1.40)	*(1.40)*	*(1.39)*	*(1.20)*	*(1.32)*	*(1.31)*	*(1.32)*	*(1.25)*	*(1.22)*	*(1.24)*
TIME	a_6	0.04039 ***	0.04035 ***	0.04045 ***	0.04239 ***	0.04045 ***	0.04053 ***	0.04054 ***	0.04111 ***	0.04143 ***	0.04131 ***
		(15.90)	*(15.88)*	*(15.89)*	*(16.47)*	*(15.89)*	*(15.91)*	*(15.92)*	*(16.06)*	*(16.15)*	*(16.12)*
Marginal cost constant	b_Q	126.58410 ***	128.83150 ***	129.04190 ***	116.56690 ***	103.64940 ***	103.22250 ***	103.96080 ***	88.03034 ***	79.42772 ***	82.65031 ***
		(10.25)	*(9.74)*	*(9.77)*	*(9.51)*	*(7.03)*	*(6.78)*	*(6.82)*	*(6.86)*	*(5.86)*	*(6.07)*
ln*Q*	b_{QQ}	-1.41624 *	-1.62944 *	-1.71909 *	-0.64992	0.80348	0.83663	0.74205	1.41688	1.94666 **	1.53998
		(-1.73)	*(-1.74)*	*(-1.81)*	*(-0.73)*	*(0.70)*	*(0.72)*	*(0.63)*	*(1.62)*	*(2.13)*	*(1.63)*
$\ln(W_1/W_2)$	b_{Q1}	19.09202 ***	19.18194 ***	19.18582 ***	23.18268 ***	19.38847 ***	19.41258 ***	19.40444 ***	16.43366 ***	15.86156 ***	15.43874 ***
		(5.13)	*(5.15)*	*(5.15)*	*(6.74)*	*(5.29)*	*(5.28)*	*(5.28)*	*(4.77)*	*(4.61)*	*(4.49)*
ln*TIME*	b_{QT}	9.25469 ***	9.57336 ***	9.39550 ***	3.28308	9.98413 ***	9.89594 ***	9.75229 ***	4.60114 *	2.92713	2.45604
		(3.31)	*(3.33)*	*(3.25)*	*(1.20)*	*(3.60)*	*(3.47)*	*(3.39)*	*(1.74)*	*(1.07)*	*(0.89)*
Lambda constant	λ_0	0.14652 ***	0.14747 ***	0.14605 ***	0.09421 **	0.12450 ***	0.12295 ***	0.11913 ***	-0.02573	-0.03537	-0.08645 *
		(3.31)	*(3.31)*	*(3.25)*	*(2.37)*	*(3.25)*	*(3.21)*	*(3.12)*	*(-1.00)*	*(-1.32)*	*(-1.70)*
MMC	λ_1	-	-0.00054	-0.00099	0.00766 **	0.00641 **	0.00638 **	0.00740 *	2.16295 ***	2.22240 **	2.96147 **
			(-0.48)	*(-0.62)*	*(2.15)*	*(2.13)*	*(2.13)*	*(1.77)*	*(2.63)*	*(2.52)*	*(2.44)*
HHI	λ_2	-	-	0.01994	0.58058 **	-	0.00325	0.05227	-	0.04671	0.61458
				(0.41)	*(2.28)*		*(0.10)*	*(0.41)*		*(1.50)*	*(1.45)*
MMC×*HHI*	λ_3	-	-	-	-0.07847 **	-	-	-0.01135	-	-	-7.76142
					(-2.29)			*(-0.40)*			*(-1.37)*
R^2 demand equation		0.9950	0.9950	0.9950	0.9952	0.9950	0.9950	0.9950	0.9951	0.9951	0.9951
R^2 cost equation		0.7128	0.7130	0.7140	0.7653	0.7223	0.7225	0.7226	0.7605	0.7650	0.7668
N.of observations		260	260	260	260	260	260	260	260	260	260
$\varepsilon_{Q,P}$		-0.26233 ***	-0.26255 ***	-0.25792 ***	-0.18979 **	-0.26531 ***	-0.26182 ***	-0.26118 ***	-0.22100 ***	-0.20988 ***	-0.21397 ***
		(-3.50)	*(-3.51)*	*(-3.44)*	*(-2.50)*	*(-3.53)*	*(-3.49)*	*(-3.49)*	*(-2.92)*	*(-2.77)*	*(-2.82)*
$\varepsilon_{Q,Z}$		0.10822 ***	0.10820 ***	0.10731 ***	0.09730 ***	0.10966 ***	0.10901 ***	0.10877 ***	0.09725 ***	0.09567 ***	0.09604 ***
		(3.24)	*(3.25)*	*(3.22)*	*(2.88)*	*(3.28)*	*(3.26)*	*(3.26)*	*(2.89)*	*(2.83)*	*(2.85)*
$\varepsilon_{P,MMC}$		-	-0.00574	-0.01067	0.00648	0.09903 **	0.09989 **	0.09983 **	0.74255 **	0.80462 *	0.79896 **
			(-0.48)	*(-0.62)*	*(0.39)*	*(2.09)*	*(2.08)*	*(2.08)*	*(2.11)*	*(1.94)*	*(1.97)*
Average estimated λ		0.14652	0.14597	0.14514	0.14103	0.15083	0.14945	0.14977	0.13776	0.13691	0.13953
Minimum estimated λ		0.14652	0.14237	0.14006	0.11217	0.13790	0.13661	0.13641	0.11458	0.11194	0.11221
Maximum estimated λ		0.14652	0.14743	0.14822	0.17362	0.17827	0.17700	0.17414	0.17560	0.17780	0.17464

The system has been estimated by non-linear three-stage least squares (*t*-values in parentheses). The demand equation includes a set of dummy variables capturing regional effects (coefficients are not reported). The instruments used are: levels and logs of first-lagged *Q* and *P*; levels and logs of *Z*, *Y*, *POP*, *BR*, average total costs, W_1, W_2, *MMC*, *HHI*, number of employees, deposits, consumption, investment, and time trend; regional dummies. The *t*-values of the elasticities are based on standard errors calculated by the delta method. Significance for the parameter estimates: *** = 1% level; ** = 5% level; * = 10% level.

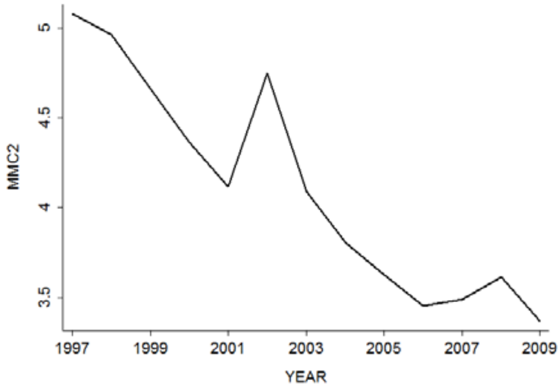
TABLE 3 – *Estimated market power indexes (fourth specification of Table 2)*

Region / Year	*1997*	*1998*	*1999*	*2000*	*2001*	*2002*	*2003*	*2004*	*2005*	*2006*	*2007*	*2008*	*2009*	*Average regional λ*
Piemonte	0.1407	0.1409	0.1409	0.1411	0.1415	0.1433	0.1445	0.1431	0.1424	0.1420	0.1432	0.1448	0.1439	*0.1425*
Valle D'Aosta	0.1588	0.1588	0.1577	0.1458	0.1470	0.1483	0.1401	0.1453	0.1458	0.1461	0.1435	0.1494	0.1457	*0.1487*
Lombardia	0.1186	0.1187	0.1189	0.1200	0.1222	0.1235	0.1230	0.1220	0.1211	0.1204	0.1240	0.1283	0.1275	*0.1222*
Trentino Alto Adige	0.1122	0.1123	0.1126	0.1127	0.1129	0.1138	0.1147	0.1147	0.1149	0.1226	0.1298	0.1221	0.1156	*0.1162*
Veneto	0.1253	0.1260	0.1268	0.1280	0.1293	0.1350	0.1399	0.1386	0.1366	0.1389	0.1418	0.1398	0.1367	*0.1341*
Friuli Venezia Giulia	0.1322	0.1311	0.1310	0.1333	0.1366	0.1402	0.1452	0.1452	0.1443	0.1450	0.1469	0.1464	0.1450	*0.1402*
Liguria	0.1507	0.1507	0.1503	0.1498	0.1499	0.1500	0.1498	0.1492	0.1486	0.1437	0.1389	0.1433	0.1496	*0.1480*
Emilia Romagna	0.1257	0.1249	0.1251	0.1258	0.1261	0.1270	0.1275	0.1269	0.1265	0.1259	0.1252	0.1263	0.1282	*0.1262*
Toscana	0.1386	0.1378	0.1371	0.1362	0.1355	0.1345	0.1341	0.1338	0.1333	0.1339	0.1341	0.1333	0.1399	*0.1355*
Umbria	0.1405	0.1403	0.1408	0.1414	0.1424	0.1415	0.1415	0.1423	0.1431	0.1432	0.1428	0.1411	0.1418	*0.1417*
Marche	0.1492	0.1479	0.1463	0.1450	0.1438	0.1427	0.1423	0.1413	0.1407	0.1397	0.1388	0.1378	0.1384	*0.1426*
Lazio	0.1457	0.1426	0.1413	0.1395	0.1383	0.1382	0.1364	0.1336	0.1328	0.1320	0.1326	0.1372	0.1397	*0.1377*
Abruzzi	0.1358	0.1363	0.1363	0.1378	0.1395	0.1389	0.1395	0.1404	0.1409	0.1415	0.1418	0.1405	0.1420	*0.1393*
Molise	0.1612	0.1601	0.1590	0.1556	0.1521	0.1519	0.1505	0.1486	0.1488	0.1494	0.1488	0.1510	0.1551	*0.1532*
Campania	0.1563	0.1546	0.1529	0.1514	0.1505	0.1529	0.1548	0.1537	0.1526	0.1515	0.1499	0.1494	0.1505	*0.1524*
Puglia	0.1308	0.1310	0.1317	0.1329	0.1351	0.1372	0.1387	0.1382	0.1381	0.1375	0.1372	0.1387	0.1430	*0.1362*
Basilicata	0.1606	0.1597	0.1587	0.1572	0.1555	0.1527	0.1493	0.1499	0.1506	0.1504	0.1495	0.1487	0.1488	*0.1532*
Calabria	0.1736	0.1707	0.1675	0.1648	0.1577	0.1524	0.1523	0.1513	0.1503	0.1499	0.1495	0.1507	0.1535	*0.1572*
Sicilia	0.1550	0.1697	0.1678	0.1625	0.1556	0.1539	0.1544	0.1529	0.1515	0.1499	0.1495	0.1516	0.1529	*0.1559*
Sardegna	0.1425	0.1385	0.1522	0.1423	0.1402	0.1366	0.1196	0.1366	0.1302	0.1364	0.1280	0.1427	0.1403	*0.1374*
Average yearly λ	*0.1427*	*0.1426*	*0.1427*	*0.1412*	*0.1406*	*0.1407*	*0.1399*	*0.1404*	*0.1396*	*0.1400*	*0.1398*	*0.1411*	*0.1419*	

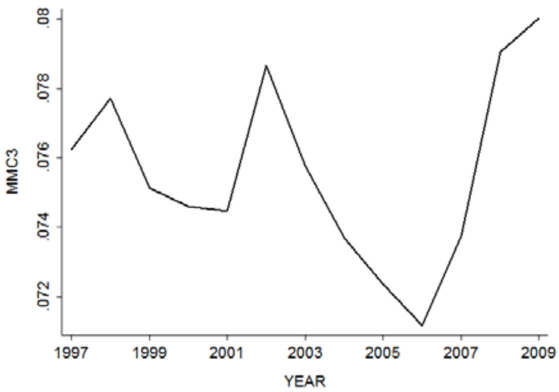
FIGURE 1 – *Multimarket contact measures and HHI by year (1997-2009) – Regional averages*



Panel (a)



Panel (b)



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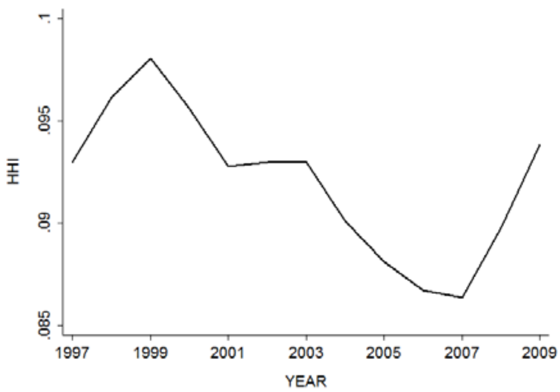
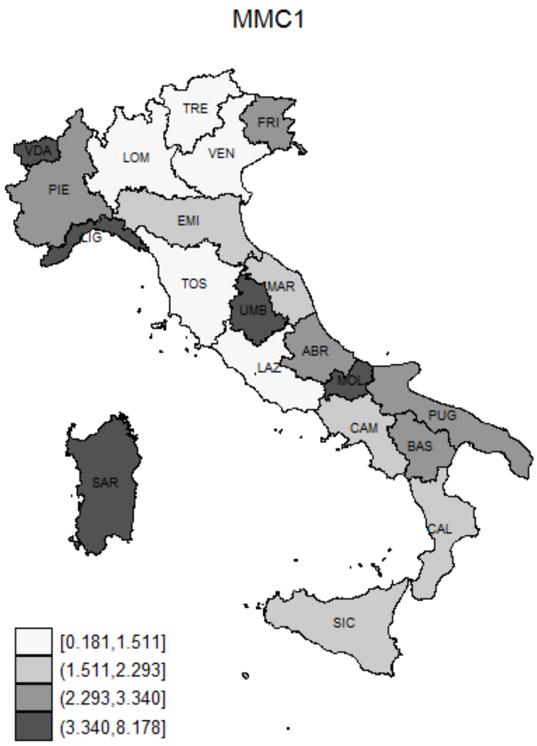
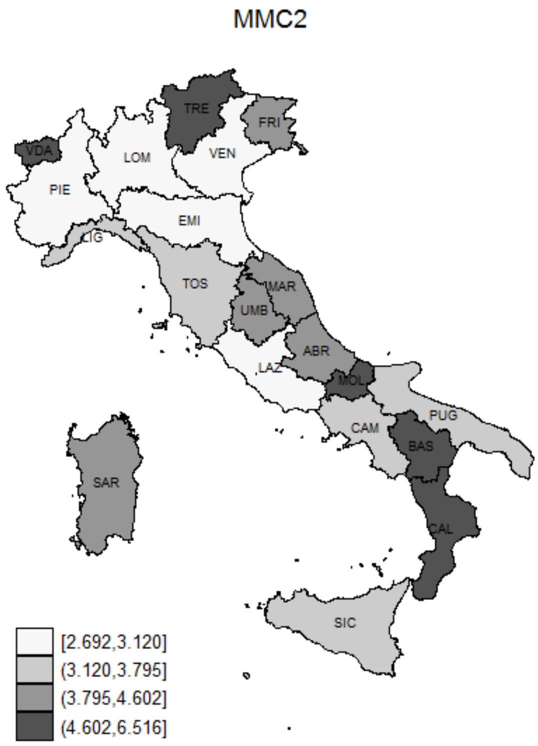


Panel (d)

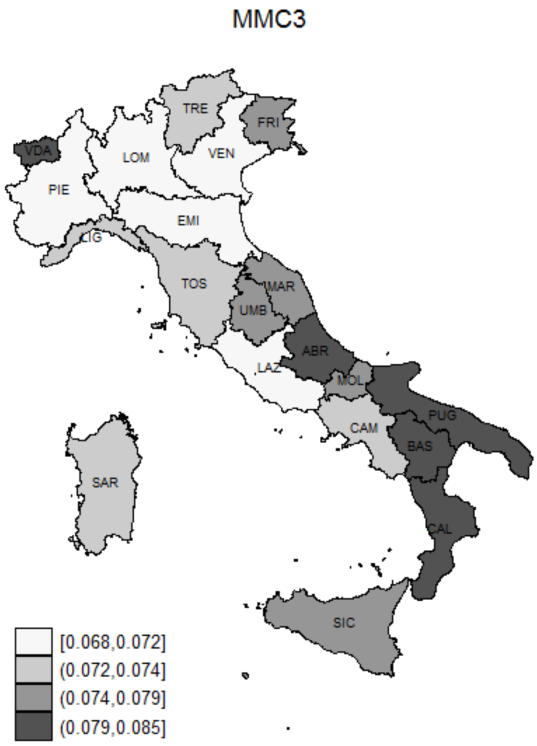
FIGURE 2 – *Multimarket contact measures and HHI by region – Yearly averages (1997-2009)*



Panel (a)



Panel (b)



Panel (c)

Panel (d)

FIGURE 3 – *Estimated regional demand, marginal cost and marginal revenues for the Italian banking industry (first specification of Table 2)*